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THE YOUTH LABOR MARKET IN THE 80's:  
DETERMINANTS OF RE-EMPLOYMENT PROBABILITIES  
FOR YOUNG MEN AND WOMEN

by

Lisa M. Lynch  
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REVISED, August 1988

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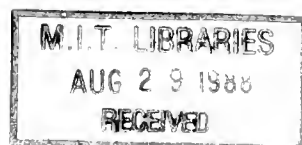
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ABSTRACT

This paper presents an analysis of the determinants of re-employment probabilities for young workers in the U.S. Using data from the new National Longitudinal Survey youth cohort a model is developed to analyze transition probabilities from nonemployment to employment. The key factors examined include personal characteristics, unemployment income, local demand conditions, and duration dependence. There are significant differences between the labor market experiences of whites and nonwhites, and males and females. High school dropouts have many more difficulties in the labor market than those who remain in school longer and/or receive other types of training. Local demand conditions are a strong determinant of the duration of spells of nonemployment and there appears to be strong evidence of negative duration dependence in re-employment probabilities for both young males and young females.

June, 1988

I would like to thank Orley Ashenfelter, Richard Blundell, Hank Farber, Larry Katz, Walter Oi and my colleagues at the National Longitudinal Survey for comments on previous drafts of this paper. I have also benefited from the comments of participants at seminars at M.I.T., Harvard, Brandeis, and the National Bureau of Economic Research and three anonymous referees. This paper is a revised version of a report prepared under contract with the Employment and Training Division, U.S. Department of Labor. Any viewpoints presented in this paper do not necessarily represent the official position and policy of the U.S. Department of Labor.



Youth unemployment in the United States, as in most other developed countries, continues to be a challenging issue for policy-makers. In particular, the recession in the early 1980's has had a profound impact on the early labor market experience of a cohort of young workers in the U.S.. The Bureau of Labor Statistics reported an overall unemployment rate in the U.S. in the fourth quarter of 1982 of 10.7 percent. However, for young workers, age 16-19, the unemployment rate was 24.3 percent, compared with 16.2 percent in 1972. For black youths the numbers were even bleaker, with 49 percent of black youths unemployed in 1982 compared with 33 percent in 1972. If one includes the number of young workers who were out of the labor force but not in school or in the military these percentages become much greater.

It might be argued that policy-makers should not be overly concerned about youth unemployment because it is simply part of a productive and efficient job search process. However, periods of unemployment (or nonemployment) may have long term employment consequences for two major reasons. First, the loss of valuable work experience may make it more difficult for youths to find employment. Labor theories such as human capital imply that since substantial investment in human capital should occur in the early years of work, early joblessness is particularly costly. If there is no investment in human capital during periods of nonemployment the entire earnings profile of the worker will be depressed.

Perhaps more importantly, dual labor market theory suggests that early nonemployment might lead to poor work habits, weak labor force attachment, and general alienation from society. The joblessness experience itself may alter the attitudes of unemployed youths if they become more discouraged about their chances of successfully finding work and this spills over into their job search behavior. In addition, employers may use employment records as a signal of potential productivity. In this context even a one time shock such as the

severe recession in 1982 may have long term consequences on the equilibrium level of unemployment, especially for young workers.

There have been numerous studies on youth unemployment in the U.S. which have used time series, cross sectional and longitudinal data. Several of these studies have focused on the factors influencing the duration of youth unemployment. Papers by Ellwood (1982), using data from the National Longitudinal Survey, NLS, young men's cohort, and Corcoran (1982), using the NLS young women's cohort, have investigated the "scarring effects" of lack of work experience in the years immediately following school leaving for U.S. youths in the late 60's and early 70's. Both studies find some evidence that after controlling for individual differences, early unemployment causes poorer future employment and earnings prospects.

Moreover, the contributions of Stephenson (1976) and (1982), Heckman and Borjas (1980), Flinn and Heckman (1982a) and (1982b), Lynch (1985), and Wolpin (1984), have tried to link theoretical developments of search theory with data on either the duration of nonemployment or unemployment of young people. The studies by Stephenson, Heckman and Borjas, and Flinn and Heckman analyzed data from the early years of the NLS young men's and women's cohorts; thus their examinations of the youth labor market focused on a period of relative economic prosperity. Utilizing data from the new NLS youth cohort allows us to examine the labor market experience of young workers in the context of a less favorable economic climate than that facing those from the earlier NLS cohorts. In particular, we shall investigate whether or not there is evidence of state dependency in youth nonemployment, and the role of personal characteristics, unemployment income, and local demand conditions in explaining the determinants of re-employment probabilities for young workers. The richer nature of the data set used, both in terms of sample size and information available on each respondent, allows us to obtain some novel results and to provide additional evidence on the importance of factors investigated in previous studies.

## I. The Model

There are different theoretical models which may be relevant for the analysis of the determinants of re-employment probabilities. The most commonly used is a simple job search model such as that presented by Mortenson (1970) and Lippman and McCall (1976). This model assumes that when a worker becomes unemployed, the expected completed duration of his or her unemployment spell (or inversely, the re-employment probability) will depend upon two probabilities -- the probability of receiving a job offer and the probability of then accepting the job offer. The probability of receiving a job offer will be determined by factors that make an individual worker more "desirable" to an employer such as education, training, ~~and~~ local demand conditions. The probability that an unemployed individual will then actually accept a job offer will be determined by his or her minimal acceptable wage or "reservation" wage. Factors which may determine this wage include the expected distribution of wages, the costs of search, any unemployment income, and the probability of receiving a job offer. The key feature to this model, which is not exclusive to job search theory, is that the waiting time to re-employment will be influenced by the probability of receiving a job offer and the reservation wage.

The re-employment probability,  $h(t)$ , which is a function of the variables described above, is also known as the hazard or failure rate in renewal theory<sup>1</sup>. The hazard rate or re-employment probability can be expressed in the following form:

$$h(t)dt = g(t)dt / (1 - G(t)) \quad (1)$$

where  $g(t)dt$  is the probability of accepting a job offer between time  $t$  and  $t+dt$ ,  $(1 - G(t))$  is the probability of not being employed at time  $t$ , and  $t$  is the duration of the current spell of nonemployment.

If we integrate eq. (1) we obtain the survivor function:

$$1 - G(t) = \exp\left[-\int_0^t h(z)dz\right] \quad (2)$$

which implies the density function:

$$g(t) = h(t)\exp\left[-\int_0^t h(z)dz\right] \quad (3)$$

It is possible to use equations (2) and (3) to develop an appropriate likelihood function for our data which will allow us to estimate the determinants of re-employment probabilities for young workers<sup>2</sup>. Our data set is composed of the stock of young workers in the NLS youth cohort not working at the 1982 interview date. We can then update their labor history on a weekly basis through the 1983 interview and observe whether or not they have been successful in finding employment during that interval. If they are successful, we can identify the exact date of their re-employment. This implies that our likelihood will contain observations on both completed and uncompleted spells of nonemployment. The appropriate form for the likelihood given this type of data is:

$$L = \prod_{i=1}^{NU} \left[ \frac{1 - G_i(t_i + h_i)}{1 - G(t_i)} \right] \prod_{j=1}^{NE} \left[ \frac{g_j(t_j + s_j)}{1 - G(t_j)} \right] \quad (4)$$

where NU is the number of individuals still jobless by the second interview,  $t$  is the duration of joblessness at the first interview,  $h$  is the number of weeks between the two interviews, NE is the number of individuals who find a job by the second interview, and  $s$  is the exact number of weeks after the first interview before becoming re-employed. This form of the likelihood allows us to control for the length bias problems associated with stock data.

In order to operationalise this likelihood function we need to select an appropriate functional form for  $G$ . Since we have censored observations (uncompleted spells of nonemployment by the second interview) we have chosen two distributions, the Weibull<sup>3</sup> and the Log-logistic, which are convenient for censored data. We assume that the hazard or re-employment probability can

be decomposed as follows:

$$h(t) = \psi_1(X)\psi_2(t) \quad (5)$$

where we assume for both distributions that:

$$\psi_1(X) = \exp[X'B] \quad (6)$$

$X'$  is a vector of appropriate explanatory variables.

For the Weibull distribution the survivor function,  $1 - G(t)$  is:

$$1 - G(t) = \exp(-\exp(X'B)t^\alpha) \quad (7)$$

and for the Log-logistic distribution the survivor function is:

$$1 - G(t) = (1 + t^\alpha \exp(X'B))^{-1} \quad (8)$$

Therefore, the hazard functions for the Weibull and the Log-logistic distributions are respectively:

$$h(t) = \alpha t^{(\alpha-1)} \exp(X'B) \quad (9)$$

$$h(t) = \alpha t^{(\alpha-1)} \exp(X'B)(1 + t^\alpha \exp(X'B))^{-1} \quad (10)$$

The Weibull hazard is monotone decreasing if  $\alpha < 1$ , monotone increasing if  $\alpha > 1$ , and constant if  $\alpha = 1$  (this is also the exponential hazard). In other words, if the hazard is constant then differences in duration spells will only be determined by the explanatory variables. If the hazard is decreasing (increasing) then, ceteris paribus, the subsequent expected duration of a spell of nonemployment will be larger(smaller) the longer the individual is not working. Job search theory predicts that as the spell of unemployment lengthens the reservation wage will fall, implying an increasing hazard (positive duration dependence). On the other hand, if employers use employment records as a signal of potential productivity or if long spells of nonemployment increase discouragement then the hazard will be decreasing (negative duration dependence).

The Log-logistic hazard is identical to the Weibull hazard apart from the term appearing in the denominator. It is monotone decreasing if  $\alpha \leq 1$  and if  $\alpha > 1$  it resembles the log-normal hazard in that it increases to a single maximum and then decreases towards zero thereafter. This formulation allows us to see

whether or not there may actually be a period of positive duration dependence followed by negative duration dependence.

## II. The Data

For our analysis we have chosen to examine the determinants of the duration of periods of nonemployment for youths in the National Longitudinal Survey, NLS, youth cohort. The NLS youth cohort is a sample of 12,686 males and females who were 14 to 21 years of age at the end of 1978. They were first interviewed in 1979 and have been interviewed every year since then about their education, jobs, military service, training programs, marriage, health, and attitudes. The response rate in 1985 was over 95 percent of the original sample. We have restricted our sample to those youths who were not employed at the 1982 interview but were not in school or in the military. In addition, we have required that the individual's date of last enrollment in school was before the starting date of their last job and that they did not return to school in the year following the 1982 interview. All of the youths in our sample have been employed immediately before their current spell of joblessness so that we restrict our analysis to the experience of young workers who appear to have entered the labor market "permanently". We did not wish to model the transition from school to first job or to include in our study those youths who only entered the labor market during the summer between school sessions. Instead, we focus on the most attached youths in the workplace. While the new NLS cohort has detailed information on starting and ending dates of employment and nonemployment spells it is not possible to observe when during a spell of nonemployment an individual is searching for work, only the proportion of time spent searching. There is some evidence using the new NLS youth cohort (Gonul (1985)) that in the 1980's there is little difference between weeks spent "unemployed" and "out of the labor force".

Using the 1983 interview we were able to determine whether or not youths not working at the 1982 interview obtained a job during the following year.



For those who were successful in finding employment we know the completed spell length of their period of nonemployment. For those still not working by the 1983 interview, we observed an "uncompleted" spell length. Approximately eighty percent of the males and sixty-three percent of the females have completed their spell of nonemployment by the 1983 interview.

Our theoretical model implies that the re-employment probability,  $h(t)$ , is determined by the probability of receiving a job offer and the probability of then accepting that job offer. Given the distribution of wages, the factors most likely to influence the probability of receiving a job offer include local demand conditions (proxied here by the local unemployment rate) and personal characteristics. We expect that nonwhites might experience longer durations of joblessness, especially if they are being discriminated against in the labor market. Other variables which might determine the probability of receiving a job offer include human capital variables such as whether or not the individual has had any training (vocational, technical, other skills, or governmental training programs) and the completed number of years of schooling. If an individual is on temporary lay-off, we would also expect that their duration of nonemployment would be shorter as long as they have a positive chance for recall. Factors affecting individuals' probability of receiving a job offer include whether or not the individual lives in an SMSA (the "inner city" problem), whether or not they are healthy, whether or not they live at home (a proxy for search intensity or parental pressure to get a job), age, and the length of the last spell of nonemployment. We also included marital status and the number of children that the respondent had.

Unemployment income subsidizes job search so that the more unemployment income an individual receives the higher the reservation wage, ceteris paribus. Many of the respondents do not receive unemployment compensation since in order to be eligible for UI a worker has to have been laid off from a previous job and to have worked for a minimum period of time. For young

workers, therefore, receipt of UI may actually send a positive signal to employers and raise the probability of receiving a job offer. Therefore, the total effect of UI on the duration of unemployment may be ambiguous. In our analysis we use the average amount of unemployment compensation the respondent receives while not working<sup>4</sup>. We tried a log value of unemployment income (assuming an extremely small value for UI for non recipients), a dummy variable for whether or not the respondent received unemployment income, and a combination of a variable equal to the log of UI for recipients and zero for non recipients and a 0-1 dummy for receipt of UI. Our final results did not alter with any of these specifications for unemployment income.<sup>5</sup>

### III. The Results

Tables 1 and 2 present our findings<sup>6</sup> on the determinants of re-employment probabilities for young males and females<sup>7</sup> and sample means of the variables. Equation 1 estimates the Weibull hazard without any explanatory variables to establish "starting" values for alpha and the log likelihood. Alpha indicates not only whether or not the hazard is increasing, decreasing, or constant but it is also a measure of misspecification (see Lancaster (1979) for a discussion of this) since it can be shown that the inclusion of significant variables will raise the value of alpha. Equations 1 and 2 assume a Weibull distribution for the hazard while equation 3 assumes the Log-logistic distribution. This allows us to examine the sensitivity of the parameter estimates to the somewhat arbitrary distributional assumptions which we have made.

Personal characteristics which lower the re-employment probability for both males and females include being nonwhite, having completed fewer years of school, and poor health. For females, being married decreases their re-employment probability<sup>8</sup>, while receiving some form of vocational or technical training, or being on layoff<sup>9</sup>, significantly raises their re-employment probability. Surprisingly, the number of children for females is not a

significant variable but this probably reflects the rather rigid selection criteria we have used to create our sample. It is interesting to observe the positive effects of private sector training for females and none for males even though column 1 of Tables 1 and 2 shows that the average amount of technical, vocational and other skills training is about the same across the two groups.

Local demand conditions seem to be critical for both males and females. This implies that these youths are not simply sampling a variety of jobs and experiencing short spells of joblessness between "samples", instead they appear to be constrained by the probability of receiving a job offer. This is also consistent with Holzer's (1986) finding that very few of those who are unemployed turn down job offers.

In an attempt to see how sensitive our parameter estimates are to the distributional assumptions we made about the hazards, we estimated both Weibull and Log-logistic hazards. While most of the coefficients remain identical (after dividing them by the value of alpha) this is not the case for the layoff coefficient in the male and female equations and the SMSA coefficient in the male equation. Being on layoff dramatically raises the probability of becoming re-employed in the Log-logistic specification compared with the Weibull specification for both males and females.

In all of the equations presented we can not reject the hypothesis that the hazard is declining and that it is declining even faster for the females than for the males. In the Log-logistic model for males, alpha equals 1.06 suggesting that the hazard first increases and then decreases. However, given the standard error we are not able to reject the hypothesis that alpha equals one and in the Log-logistic case this implies a declining hazard. This suggests that for both males and females the probability of becoming re-employed, ceteris paribus declines with the duration of a spell of nonemployment, contrary to what a simple job search model would predict.

Since the coefficients presented in Tables 1 and 2 are not particularly

intuitive<sup>10</sup> we have calculated the expected completed duration of a spell of nonemployment for different types of individuals using the coefficients from equation 2 in Tables 1 and 2. For example, integrating the Weibull survivor function, equation 7, we obtained the expression for the expected completed duration of a spell of nonemployment:

$$E(T) = \Gamma(1/a + 1) \exp(-(X'B)/a) \quad (11)$$

The first expected duration calculated was for a "typical" male. A "typical" male is white, 21 years old in 1982, has completed 11 years of school, unmarried with no children, lives in an SMSA with an 11 percent unemployment rate, healthy, living at home, has no formal job training, and has had a past spell of nonemployment of 15 weeks. The expected completed duration of joblessness for this typical male is 7.2 weeks. If we make this male nonwhite his expected spell length increases to 14.1 weeks. If instead he lives in an SMSA with a 20 percent unemployment rate his spell increases from 7.2 weeks to 19 weeks. Finally, if this male is "typical" but he has finished college his expected completed duration is only 2.6 weeks.

If we repeat this exercise for females (same characteristics as for males except that completed years of school equals 12 and the length of the past spell of nonemployment equals 18 weeks) we find that the expected completed duration of joblessness for a typical female is only 5.4 weeks. However, if we make this female nonwhite her spell length increases dramatically to 24 weeks. If instead she lives in an SMSA with a 20 percent unemployment rate her spell length increases from 5.4 weeks to 17.8 weeks. If she has completed college her expected spell is only 1.2 weeks or if she has taken some sort of vocational, technical or other skills training her expected spell length is only 2.6 weeks. Finally, if she has all of the "typical" characteristics except that she is married and is not living at home with her parents her expected completed duration of nonemployment increases to 20.3 weeks. This last finding highlights an important difference between the males and

females in our sample. However, this effect does not seem to be the result of children since that variable is never significant in our estimation.

Using equation 9, the re-employment probability for either a "typical" male or female who has not been working for one week is slightly greater than thirty percent. If they have not been working for eight weeks this drops to eight percent and if they have not been employed for fifty two weeks their re-employment probability is only two percent.

#### IV. Unobserved Heterogeneity

Before concluding that there is negative duration dependence in the transition probability from nonemployment to employment it may be possible that our parameter estimates are biased due to the omission of unobserved variables such as motivation. As is well known this may lead to spurious negative duration dependence. Equation 1 in Tables 1 and 2 presents estimates of alpha excluding all of the explanatory variables for males and females. Including a wide range of observable factors only increases the estimate of alpha from .32 to .37 for the males and from .2 to .3 for the females. It is unlikely that any unobserved heterogeneity which remains in our sample will have an effect large enough on alpha to raise it significantly over one. Nevertheless, in this section we present some alternative approaches to take into account the impact of unobserved heterogeneity on our parameter estimates.

We first attempted to control for unobserved heterogeneity using the Gamma Weibull mixing distribution (see Lancaster (1979)). The model kept collapsing to the Weibull case so other semi-parametric or non-parametric techniques were explored. We then estimated Cox's (1972) proportional hazard model. This model is nonparametric in the sense that it involves an unspecified function in the form of an arbitrary base-line hazard function. While using this method means that we will can not examine state dependency in youth nonemployment it does allow us to examine the robustness of the estimates of other parameters of interest. While the coefficients obtained using Cox's proportional hazard

model are somewhat smaller than those obtained assuming a Weibull distribution, they do not appear to be significantly different.<sup>11</sup>

Another approach is that of Heckman and Singer (1984) who control for unobserved heterogeneity in a nonparametric way. Assume that we have a population divided into  $J$  homogeneous groups. The hazard then for individual  $i$  with measured covariates  $X_i$  in group  $j$  is:

$$h_{ij}(t) = \alpha t^{(\alpha-1)} \exp(X_i' B + \lambda_j) \quad (13)$$

The number of groups,  $J$ , is unknown so we began with  $J=1$  and incremented until the likelihood failed to show significant improvement. The results presented in Tables 1 and 2 are equivalent to the case of  $J=1$ . At  $J=2$  there was no significant improvement in the value of the likelihood and all of the parameter estimates were virtually identical to those obtained with  $J=1$  for both males and females. However, as Trussell and Richards (1986) have shown, even with this type of nonparametric representation of heterogeneity, results can still be sensitive to the choice of the hazard. In addition, the cutoff point for  $J$  is always arbitrary in the sense that while  $J=2$  or  $J=3$  may not alter the likelihood, higher values for  $J$  may. Apart from creating a truly nonparametric estimator, all we can do is to try a variety of approaches and observe what happens to our estimates. In our case, the fundamental finding is that the results presented in section 3 appear to be quite robust even when the heterogeneity problem is accounted for. This may reflect the relatively greater homogeneity of our sample compared with other samples that have been analyzed. In addition, given the small percentage of our sample that receives unemployment income, we did not encounter the "26 week spike in the hazard" problem that other investigations have faced.

## V. Conclusions

This paper has tried to identify significant factors which affect the length of joblessness for youths. Few studies have examined racial or sexual differences in re-employment probabilities. One exception is Ellwood (1982)

who did examine the impact of race on male re-employment probabilities but found it insignificant for the late sixties and early seventies. We find that the re-employment probability for both young males and females in our sample is significantly reduced if the individual is nonwhite, even after controlling for a wide range of other characteristics. In spite of affirmative action legislation nonwhites still seem to be experiencing much more difficulty than their white counterparts in the labor market. Our results may even underestimate the "true" racial differences among young workers since we have conditioned our estimation on having been employed before becoming nonemployed.

Investments in human capital are important determinants of the probability of transiting from nonemployment to employment. There are high returns schooling for both the males and females. Unfortunately, many of the youths in our sample (forty percent of the females and fifty two percent of the males) have not even completed high school. We also find that private sector training such as vocational, technical or on-the-job training significantly increases the re-employment probability of females. However, this variable does not seem to affect males even though the same percentage of males and females have had this type of training. Training obtained in some sort of government program does not seem to be very effective in improving the chances of these youths to become re-employed.

There are also other significant differences in the labor market experience of male and female workers. There is evidence that the non labor market activities of females alters their labor market experience. In particular, as females become married their nonemployment spells lengthen significantly. However, this does not seem to be due to the presence of children since this variable is never significant.

Perhaps as important as the variables that are significant are those that are not. In particular, unemployment compensation is never a significant explanatory variable in any of our equations. We would not expect, therefore,

that cuts in unemployment compensation would be an effective tool to shorten the length of unemployment spells for young workers.

Previous studies such as Ellwood (1982) have not always found a significant effect of local demand on spells of unemployment, especially during periods of relative economic prosperity. However, our analysis of a period covering a severe recession, demand conditions play an important role for both males and females. Being in a depressed area more than doubles the expected duration of nonemployment for males and more than triples the spell length for females. Even though we have controlled for local unemployment conditions it may be possible that some of our results are influenced by a worsening of economic conditions over the observation period. However, our data cover a period where the unemployment rate for 16-24 year olds was high but relatively constant.

Finally, Heckman and Borjas (1980) and Flinn and Heckman (1982) have examined the issue of duration dependence with a small sample from the earlier NLS young men's cohort. Heckman and Borjas conclude on the basis of regression models that there is no negative duration dependence in the completed duration of unemployment spells. However, using the same data set and maximum likelihood estimation, Flinn and Heckman (1982) reach the opposite conclusion. All of these authors express concern over the limited sample size used. With our larger and more detailed data set we find that there is strong evidence of negative duration dependence in re-employment probabilities for young workers in the 80's. This implies that as the spell length of nonemployment increases the probability of becoming re-employed declines sharply. This may be the result of employers using the length of a youth's spell of nonemployment as a signal of some undesirable but unobservable characteristic. It may also reflect the fact that six months of joblessness for a young worker may have a very different impact on their attitudes towards the workplace than six months of unemployment to an adult worker with twenty years or more of employment.



## FOOTNOTES

- 1.) See Cox (1962), Kalbfleisch and Prentice (1980), and Cox and Oakes (1984).
- 2.) We estimate a reduced form model so we observe the total effect of variables on re-employment probabilities rather than the separate effects on the reservation wage and the probability of receiving a job offer. Therefore our estimation is not a direct test of search theory, but it does have the advantage of not imposing the tight distributional assumptions required in structural analyses. From a policy point of view it seemed that the total effects of various factors on re-employment probabilities were of greater interest.
- 3.) See Lancaster (1979) for a complete discussion on the Weibull hazard.
- 4.) Average unemployment compensation is the average benefit received while eligible for UI.
- 5.) Previous studies have included the ratio of unemployment income to expected earnings. This "replacement ratio" has no specific justification within a job search framework so in the technical appendix available from the author we include the two variables separately in our estimation. Even using a fitted value from an earnings function as a proxy for expected earnings results in problems of enogeneity (see Nickell (1979)) so we present results of a true reduced form model where we do not include expected earnings.
- 6.) The estimates presented here were obtained using a modified Newton algorithm from the National Algorithms Group Library, number E04LBF which required analytically derived first and second derivatives. These derivatives were checked numerically with NAG routines E04HCF and E04HDF.
- 7.) We have analyzed the re-employment probabilities for males and females separately since the restriction of combining the two groups was rejected by a likelihood ratio test.
- 8.) The criteria used for creating the sample results in only 18 out of 761 males "out of the labor force" and 136 out of 892 females "OLF" over the sample period. Given concerns raised in Flinn and Heckman (1983) about the distinction between OLF and unemployment all equations were re-estimated deleting the OLF observations. There was no change in the results for males and for the sample of "unemployed" females all of the coefficients remained the same with the exception of marital status which became insignificant.
- 9.) It may be possible that those who have been on layoff less than 3 months may actually have rising hazards. Analysis of this sub sample was not possible since the sample sizes were too small (less than 40 for the girls and less than 56 for the boys).
- 10.) For the Weibull  $d\log E(T)/dX_i = -B_i \lambda$ . See Lancaster (1979) for a detailed discussion of this.
- 11.) Results available in the technical appendix. It should be noted that the asymptotic normality of our estimates given this likelihood has not yet been formally proven. See Cox and Oakes (1984).

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Table 1 Maximum likelihood estimates of re-employment probabilities  
(standard errors in parentheses)  
Males (N = 761)

| Variable                   | Mean<br>(Std. dev.)  | Eq. 1<br>(Weibull) | Eq. 2<br>(Weibull) | Eq. 3<br>(Log-logistic) |
|----------------------------|----------------------|--------------------|--------------------|-------------------------|
| CONSTANT                   |                      | 0.02<br>(0.04)     | -0.10**<br>(0.04)  | 7.50**<br>(0.23)        |
| ALPHA                      |                      | 0.32*<br>(0.04)    | 0.37*<br>(0.04)    | 1.06**<br>(0.15)        |
| NONWHITE                   | .36<br>(.48)         | -                  | -0.23**<br>(0.09)  | -0.69**<br>(0.34)       |
| AGE                        | 21<br>(2.13)         | -                  | -0.02<br>(0.02)    | -0.02<br>(0.09)         |
| SCHOOL                     | 11.07<br>(1.65)      | -                  | 0.045**<br>(0.027) | 0.15<br>(0.12)          |
| MARRIED                    | .14<br>(.35)         | -                  | 0.16<br>(0.12)     | 0.89<br>(0.86)          |
| LOG UB                     | .21<br>(.41)         | -                  | 0.01<br>(0.01)     | 0.01<br>(0.05)          |
| CHILDREN                   | .08                  | -                  | -0.12<br>(0.08)    | -0.37<br>(0.28)         |
| LOCAL URATE                | 10.73<br>(3.64)      | -                  | -0.04**<br>(0.01)  | -0.14**<br>(0.04)       |
| HEALTHY                    | .95<br>(.23)         | -                  | 0.46**<br>(0.21)   | 1.31**<br>(0.66)        |
| LIVE AT HOME               | .63<br>(.48)         | -                  | 0.14<br>(0.10)     | 0.40<br>(0.42)          |
| TRAINING VTS               | .28<br>(.45)         | -                  | -0.10<br>(0.09)    | -0.42<br>(0.44)         |
| TRAINING GOV               | .07<br>(.25)         | -                  | 0.01<br>(0.17)     | -0.03<br>(0.70)         |
| LAYOFF                     | .14<br>(.34)         | -                  | 0.23<br>(0.13)     | 11.62**<br>(1.34)       |
| SMSA                       | .70<br>(.46)         | -                  | -0.16**<br>(0.09)  | -10.72**<br>(3.68)      |
| LAG DURATION<br>DEPENDENCE | 15.36 wks<br>(13.66) |                    | -0.005<br>(0.003)  | -0.01<br>(0.01)         |
| LOG LIKELIHOOD =           |                      | -2476.94           | -2454.37           | -2470.41                |

\*Tests whether the value is significantly different from one, \*\* indicates significantly different from zero.

Table 2 Maximum likelihood estimates of re-employment probabilities  
(standard errors in parentheses)  
Females (N = 892)

| Variable                   | Means<br>(Std. Dev.)   | Eq. 1<br>(Weibull) | Eq. 2<br>(Weibull) | Eq. 3<br>(Log-logistic) |
|----------------------------|------------------------|--------------------|--------------------|-------------------------|
| CONSTANT                   |                        | 0.43**<br>(0.04)   | -1.13**<br>(0.04)  | -4.10**<br>(0.12)       |
| ALPHA*                     |                        | 0.20**<br>(0.04)   | 0.30**<br>(0.04)   | 0.87**<br>(0.04)        |
| NONWHITE                   | .32<br>(.47)           | -                  | -0.45**<br>(0.10)  | -1.29**<br>(0.25)       |
| AGE                        | 21<br>(2.13)           | -                  | 0.001<br>(0.02)    | 0.03<br>(0.05)          |
| SCHOOL                     | 11.46<br>(1.60)        | -                  | 0.11**<br>(0.03)   | 0.30**<br>(0.07)        |
| MARRIED                    | .41<br>(.49)           | -                  | -0.25**<br>(0.97)  | -0.38<br>(0.25)         |
| LOG UB                     | .10<br>(.29)           | -                  | 0.02<br>(0.02)     | 0.05<br>(0.05)          |
| CHILDREN                   | .51<br>(.50)           | -                  | 0.002<br>(0.06)    | 0.25<br>(0.16)          |
| LOCAL URATE                | 10.46<br>(3.62)        | -                  | -0.04**<br>(0.01)  | -0.16**<br>(0.03)       |
| HEALTHY                    | .86<br>(.35)           | -                  | 0.36**<br>(0.14)   | 0.85**<br>(0.32)        |
| LIVE AT HOME               | .38<br>(.49)           | -                  | 0.15<br>(0.11)     | 0.53<br>(0.30)          |
| TRAINING VTS               | .30<br>(.46)           | -                  | 0.22**<br>(0.09)   | 0.63**<br>(0.30)        |
| TRAINING GOV               | .06<br>(.24)           | -                  | -0.05<br>(0.19)    | 0.26<br>(0.55)          |
| LAYOFF                     | .05<br>(.22)           | -                  | 0.66**<br>(0.16)   | 7.11**<br>(1.46)        |
| SMSA                       | .71<br>(.45)           | -                  | -0.10<br>(0.10)    | -0.29<br>(0.26)         |
| LAG DURATION<br>DEPENDENCE | 17.74 weeks<br>(14.76) | -                  | -0.003<br>(0.003)  | -0.01<br>(0.01)         |
| LOG LIKELIHOOD =           |                        | -2486.63           | -2444.74           | -2459.97                |

\*Tests whether the value is significantly different from one, \*\* indicates significantly different from zero.











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